APPENDIX ON: "WELFARE EFFECTS OF SWITCHING BARRIERS THROUGH PERMANENCE CLAUSES: EVIDENCE FROM THE MOBILES MARKET IN COLOMBIA"

1 Introduction

In this appendix we present a robustness check for the results in the study "Welfare effects of switching barriers through permanence clauses: Evidence from the mobiles market in Colombia". We describe the benefits of estimating only the demand side of the market, the details of the estimation algorithm, the characteristics of a more recent set of data, and the counterfactual experiments. The underlying results show no substantial difference with those presented in the original study. This additional empirical evidence supports the conclusions outlined in the study. Recall the goal of the original study is to measure the effects on social welfare associated with the enactment of Resolution 4444, by which the *Comisión de Regulación de Comunicaciones* banned the fixed-length terms in mobile telecommunications contracts also known as *permanence clauses*.

2 Model

To measure the effects on social welfare of the restriction on permanence clauses we adopt the methodology introduced by Berry et al. (1995) to develop a structural model of demand for mobile terminals. Consumer i in market t makes discrete choices over mobile terminals based on the indirect utility level, u_{ijt}, she perceives from the permanence clause c_{jt}, a set of K observable characteristics, $\mathbf{x}_{jt}$, and an unobservable (to the econometrician) attribute, ξ_{jt}, associated to terminal j, as well as from her income, y_{it}, and a random shock to her utility ε_{ijt}:

$$u_{ijt} = \alpha_i(y_{it} - p_{jt}) + \mathbf{x}_{jt}\beta_i + \tau_i c_{jt} + \xi_{jt} + \varepsilon_{ijt}. \tag{1}$$

Consumer heterogeneity is captured both through differences in y_{it} and differences in β_i. The consumer-specific parameters of the indirect utility function (1) can be defined as follows:

$$\begin{pmatrix} \alpha_i \\ \beta_i \\ \tau_i \end{pmatrix} = \begin{pmatrix} \overline{\alpha} \\ \overline{\beta} \\ \overline{\tau} \end{pmatrix} + (\Sigma v_i + \Pi y_{it}), \tag{2}$$

where Σ is a $(K+1) \times (K+1)$ scaling matrix of coefficients, and Π is a $(K+2) \times 1$ vector of coefficients.

We can rewrite the indirect utility function as:[1]

$$u_{ijt} = \alpha_i y_{it} + \delta_{jt} + \mu_{ijt}, \tag{3}$$

where δ_{jt} denotes the average utility and μ_{ijt} represents the consumer-specific deviations from the average utility, as shown below:

$$\delta_{jt} = \mathbf{x}_{jt}\overline{\beta} - \overline{\alpha} p_{jt} + \overline{\tau} c_{jt} + \xi_{jt},$$
$$\mu_{ijt} = [-p_{jt}, \mathbf{x}_{jt}, c_{jt}](\Sigma v_i + \Pi y_{it}).$$

The *outside option*, which the consumers are also able to choose, is denoted by alternative $j = 0$. The indirect utility level for the outside alternative is given by:

$$u_{i0t} = \alpha_i y_{it} + \xi_{0t} + \sigma_0 v_{i0} + \varepsilon_{i0t}. \tag{4}$$

Following the convention in these models, we normalize $\delta_{0t} = 0$, since this term cannot be identified given the methodology and available data.

The consumer i chooses the mobile j that generates the highest utility level compared to all other available options, $u_{ijt} \geq u_{ikt}$. Assuming that ε follows an extreme value type I distribution, choice probabilities for consumer i can be expressed in a closed-form fashion as:

$$s_{ijt}(p_{jt}, \mathbf{x}_{jt}, \delta_{jt}, P_{ns}; \theta) = \frac{e^{\delta_{jt}+\mu_{ijt}}}{1 + \sum_{k=1}^{J} e^{\delta_{kt}+\mu_{ikt}}}, \tag{5}$$

[1] We keep the term $\alpha_i y_{it}$ for consistence with the previous equation but, as pointed out in the original study, this term vanishes from the utility specification

with product aggregate market shares:

$$s_{jt}(p_{jt}, \mathbf{x}_{jt}, \delta_{jt}, P_{ns}; \theta) = \int_{A_{jt}} s_{ijt} dP_o(v) dP_o(y), \tag{6}$$

where $A_{jt} = \{(y_{it}, v_i, \varepsilon_{ijt}) | u_{ijt} > u_{ikt}, \forall k \neq j\}$ is the set of consumer traits that lead to the purchase of mobile j.

Assuming that the empirical distribution of income is Log-Normal, where $v_y \sim N(0,1)$, and $v = (v_{i1}, ..., v_{ik}, v_{iy})$ are independent variables drawn from $P_o(v)$, we can reduce the integral above to one dimension as:

$$s_{jt}(p_{jt}, \mathbf{x}_{jt}, \delta_{jt}, P_{ns}; \theta) = \int_{A_{jt}} s_{ijt} dP_o(v). \tag{7}$$

And then approximate the market share through an average over ns random draws from $P_o(v)$ so that $\lim_{ns \to \infty} P_{ns}(v) \to P_o(v)$:

$$s_{jt}(p_{jt}, \mathbf{x}_{jt}, \delta_{jt}, P_{ns}; \theta) = \frac{1}{ns} \sum_{i=1}^{ns} s_{ijt}. \tag{8}$$

The price derivatives of demand are given by equation (9). The first line of the equation shows the own-price derivatives and the second line shows the derivative of market shares with respect to all other prices.

$$\frac{\partial s_{jt}}{\partial p_{jt}} = -\int_{A_j} \alpha_i s_{ijt}(1 - s_{ijt}) dP(v), \tag{9}$$

$$\frac{\partial s_{jt}}{\partial p_{kt}} = \int_{A_j} \alpha_i s_{ijt} s_{ikt} dP(v). \tag{10}$$

On the supply side assume there are F firms in the market, each producing a subset $\mathcal{F}_f$ of mobile terminals. Producers compete via prices, so that the Nash-Bertrand equilibrium is given by the set of equations that solve the problem of profit maximization simultaneously. The profits function is:

$$\pi_{ft} = \sum_{j \in F_f} (p_{jt} - mc_{jt}) M s_{jt}, \tag{11}$$

where M is the size of the potential market and mc_{jt} is the marginal cost associated with product j in market t. The first order conditions for profit maximization with respect to the price of terminal j is given by:

$$s_{jt} + \sum_{r \in F_f} (p_{rt} - mc_{rt}) \frac{\partial s_{rt}}{\partial p_{jt}} = 0 \tag{12}$$

Let Ω be a matrix with derivatives of market shares with respect to prices, such that:

$$\Omega_{jr} = \left\{ \begin{array}{ll} -\frac{\partial s_{rt}}{\partial p_{jt}} & \text{if } f \text{ produces both } r \text{ and } j \\ 0 & \text{o.w} \end{array} \right\}. \tag{13}$$

Hence, rewriting the first order conditions in matrix form and rearranging terms, we obtain an expression for price markups, which depends only on demand parameters: price derivatives and market shares:

$$b = p - mc = \Omega^{-1} s. \tag{14}$$

Note the model allows us to recover marginal costs without any knowledge on firm technology. That is, with the use of proper instruments for the potential correlation between ξ_j and the price equations, we can estimate the model without specifying a marginal costs equation. Moreover, this alternative leads to more efficient estimators as there are less parameters to estimate with the same data set.

3 Estimation

As we mentioned before, demand parameters suffice to perform our two counterfactual scenarios. That is, $\theta_1 = (\overline{\alpha}, \overline{\beta}, \overline{\tau})$, as well as $\theta_2 = (\Sigma, \Pi)$ and ξ_{jt}. The estimation process follows the methodology introduced by Berry et al. (1995). We begin by setting an initial point for $\theta_2 = \theta_2'$ in the respective parameter space. Conditional on a fixed point in the space of θ_2, δ_{jt} can be solved numerically using a contraction mapping:

$$s_{jt}(p_{jt}, \mathbf{x}_{jt}, \delta_{jt}, P_{ns}; \theta) = s_{jt}, \tag{15}$$

which amounts to computing:

$$\delta_{jt}^{(n+1)} = \ln(s_{jt}) - \ln(s_{jt}(p_{jt}, \mathbf{x}_{jt}, \delta_{jt}, P_{ns}; \theta)) + \delta_{jt}^{(n)}, \tag{16}$$

for each $n = 1, 2, \ldots,$ iteration. δ_{jt} is the updated in each iteration by the difference between the log of observed market shares, s_{jt}, and the log of model-predicted market shares, $s_{jt}(p_{jt}, \mathbf{x}_{jt}, \delta_{jt}, P_{ns}; \theta)$ until convergence is reached.

Once $\delta_{jt}(\theta_2')$ is solved for, we proceed to estimate θ_1. Then, we estimate $(\overline{\alpha}, \overline{\beta}, \overline{\tau})$ by regressing the respective mean utility on the observed product characteristics and price instruments.

Using the structural estimates of the demand unobservable, ξ_{jt}, the non-linear parameters, θ_2, can be computed by the following GMM estimator:

$$\theta_2^* = \arg\min_{\theta_2} \ \ \xi(\theta_2)' Z W Z' \xi(\theta_2), \tag{17}$$

where Z is the matrix of instruments and W is a weighting identity matrix. The estimator for the variance-covariance matrix is computed as suggested by Cameron and Trivedi (2005).

4 Counterfactuals

Switching costs by means of permanence clauses are measured by the parameter τ in the model which interacts with an indicator variable of mobile terminals sold with the clause. This indicator variable takes the value of 1 for terminals with postpay calling plans sold during the months previous to July 2014. We expect its sign to be negative suggesting permanence clauses reduce consumer utility.

In addition to the raw estimate of switching costs using τ, we model changes in overall consumer welfare and firm profits by estimating two counterfactual scenarios: one in which we assume Resolution 4444 was never enacted, hence permanence clauses would still be permitted and the indicator variable would take the value of 1 for all terminals sold with postpay calling plans during the entire time database; and another in which we assume permanence clauses did not exist from the beginning of our sample (January 2014), hence the indicator variable would be zero always.

Formally, let the upper script C denote the counterfactual scenario and the upper script O denote the observed scenario, equilibrium prices in the counterfactual solve the following equation:

$$p^C = mc^O + \Omega(p^C)^{-1} s(p^C, x^C, \xi, P_{ns}; \theta). \tag{18}$$

Equation (18) relies in several important assumptions which are discussed in the main study. In light of our data, these assumptions are reasonable given that prices reflect only the price of the handset but not the price of the hired mobile services.

After computation of the market equilibrium under each counterfactual scenario, changes in consumer welfare under the BLP methodology can be obtained from equation (19), where $V_{ijt} = \delta_{jt} + \mu_{ijt}$. On the supply side, changes in firm surplus can be computed from equation (20), and the sum of these two expressions gives us the variation in welfare from the societal perspective (equation (21)).

$$\Delta C = M \int \frac{log(\sum_{j=1}^{J} e^{V_{ijt}^O}) - log(\sum_{j=1}^{J} e^{V_{ijt}^C})}{\alpha_i} dP(v) \tag{19}$$

$$\Delta E = \sum_f \pi_f^O - \pi_f^C = \sum_f \left(\sum_{j \in \mathcal{F}_f} (p_{jt}^O - cmg_{jt}^O) M s_{jt}^O - \sum_{j \in \mathcal{F}_f} (p_{jt}^C - cmg_{jt}^C) M s_{jt}^C \right) \tag{20}$$

$$\Delta S = \Delta C + \Delta E \tag{21}$$

5 Data

The data used in the main study to estimate the demand for mobile terminals in Colombia and the effect of permanence clauses consisted of monthly series for a sample of mobiles sold between January 2014 and June 2016. The results shown in this appendix are generated using the same data plus a more recent set comprising information from July 2016 to June 2017. Both data sets were collected by GfK, a market research company, and was given to us by Fenalco, an alliance of retail sales companies.

The data includes sales by retailers who authorized GfK (with written acceptance) to share the information with Fenalco. An observation in our data set is identified by the combination of mobile reference, tariff plan (postpay, prepay, SIM free), distribution channel (authorized distributors, AD, and department stores, DS) and month. For each mobile reference we observe units sold, price, screen size, and memory size in GB. For a single mobile reference, price varies only over operators, distribution channels, and time. The price variable was build by GfK in collaboration with network operators. It corresponds to the individual price of each mobile after removing the over-pricing due to voice and data plans. We define a market as a month and use the distribution channel as a product observable characteristic since a mobile reference can compete with itself in a market if being sold through different channels. We identify 5 producers in each market: Claro, Movistar, Tigo, SIM free in authorized distributors and SIM free in department stores. Since we can not observe all the individual SIM free producers, we assume those selling through authorized distributors act as a single producer and those selling in department stores act as another single producer.[2] Claro, Movistar and Tigo are the only ones who can offer bundled sales of terminals and plans because they are also the only network operators in the market. SIM free producers, as their name suggests, only sells terminals at cash price without a calling plan. In terms of consumer characteristics, the mean and standard deviation of the income distribution are obtained from the GEIH survey of the National Administrative Department of

[2] Given how the data is constructed, there is natural concern about potential selection biases. The retailers that authorized GfK are usually those with the highest number of sales, a characteristic that is evidently correlated with the unobserved terms in our model. A proper Heckman selection model to avoid such biases would require to specify a model that describes the probability of authorizing GfK to use their sales data as a function of a set of observed characteristics of the retailers themselves or of the terminals they sell. However, in the Colombian market there is no other available source with this sort of data for such small retailers. Therefore, although we acknowledge the potential for selection bias, we believe this set of retailers to be small enough not to influence the results. Moreover, this belief is supported by the similar estimations results we obtain for the same model using a more recent set of data, which includes one more year of information. These results can be seen in an available online appendix `http://www.alvaroriascos.com/`

Statistics. These parameters are allowed to vary each month. We define the potential market as the employed population.

Figure 1 shows the aggregate quantity of handsets sold in the market. The black vertical line marks the month when Resolution 4444 was officially enacted.

Figure 1: Evolution of monthly aggregate sales in the market

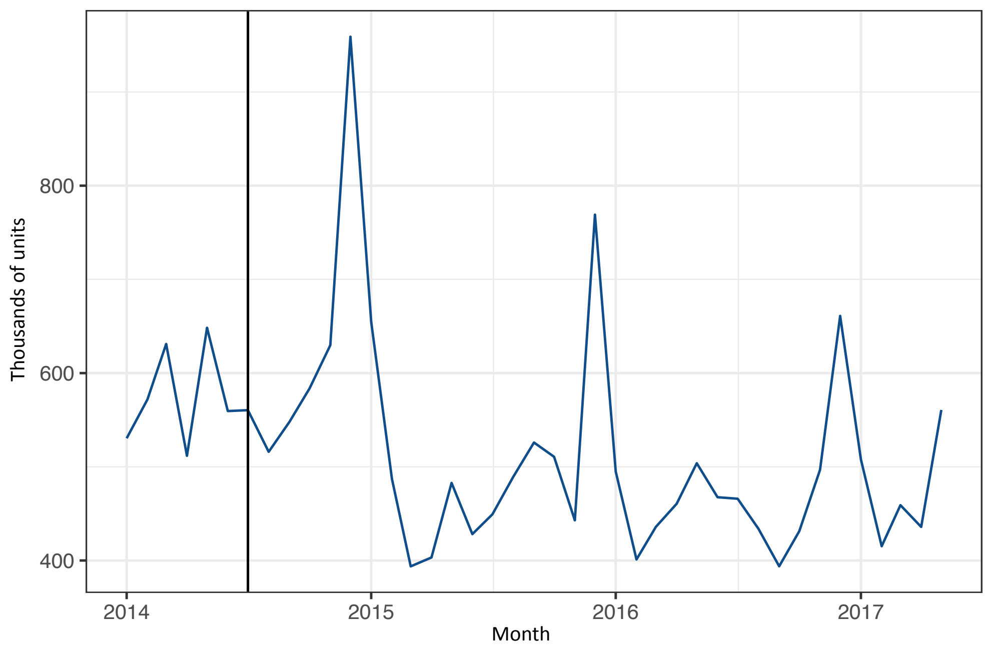
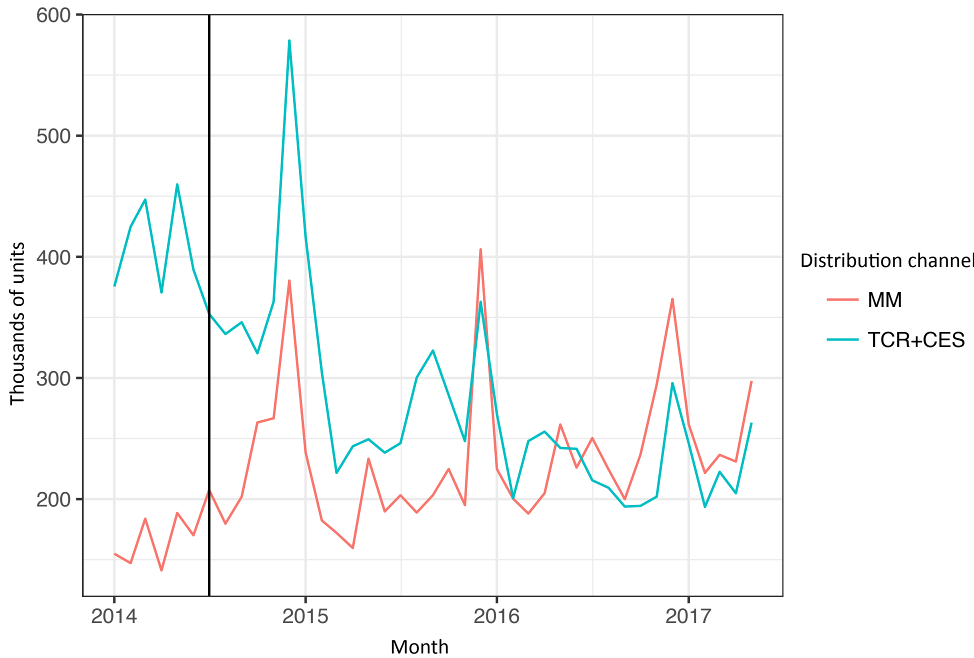


Figure 2 presents the aggregate quantity of handsets sold by each type of retailer, that is mobile services providers' stores (TCR+CES) and independent retailers (MM). As it can be observed MM retailers start to gain the majority of the market by 2017. This change in the market shares suggests that it is worth to explore whether the conclusions made in the main study still hold with the addition of the new data.

Figure 2: Evolution of monthly sales by type of retailer



6 Results

Below we report the estimation results of the model using the most recent database and focusing in estimating only the demand-side parameters. Interactions on product observable characteristics include: permanence clause, screen size, and memory size. We use the Nelder-Mead simplex routine to minimize the GMM function.

The estimated parameters and associated standard errors are presented in Table 1. The first column reports mean marginal utilities ($\overline{\alpha}, \overline{\beta}, \overline{\tau}$), the second column deviations from the mean (Σ), and the third column income interactions terms (Π).

The results shown in this table are very similar to the ones presented in the main study, regardless of the addition of new data and the modification in the estimation algorithm. All the coefficients of the mean marginal utilities are significantly different from zero. Larger screen size is associated on average to higher utility as well as a larger memory size. The coefficient on the permanence clause indicator is negative, as we expected, and suggests consumers experience lower utility from buying handsets associated with a permanence clause in their mobile service contract. In other words, the coefficient measures the differential cost in utility due to the permanence clause.

Table 1: Estimated coefficients of the demand equations

Variable	Estimate	Standard deviation	Interaction with income
Constant	-103.315^{***}	3.313^{***}	1.889^{***}
	(0.012)	(0.389)	(0.059)
Permanence clause	-8.926^{**}	1.222^{***}	1.660^{***}
	(0.002)	(0.171)	(0.014)
Price	-31.667^{***}	1.659^{***}	1.580^{***}
	(0.007)	(0.022)	(0.015)
Memory size	22.446^{***}	2.389^{***}	1.306^{***}
	(0.005)	(0.054)	(0.020)
Screen size	25.034^{***}	0.356^{***}	1.306^{***}
	(0.008)	(0.138)	(0.006)

Note: The first column shows the mean marginal utility, the second column shows the deviations from the mean, and the third the interactions with income. Standard errors are in parenthesis. Authors' calculations based on data collected by GfK.

The results regarding standard deviations and interactions with income are also very similar to those shown in the original study. For the permanence clause, in particular, we find that, despite its estimated volatility across consumers, consumers will substitute away disproportionately towards cash price mobiles (with or without calling plans) when the price of a terminal with permanence clause increases. Notice that, even if we add two standard deviations to the average marginal utility generated by the permanence clause, the effect of permanence clause is still negative.[3] This means that is very unlikely to find consumers who may derive higher utility from being locked-in than from having the option of buying the terminal and the plan separately.

To derive a measure of the switching cost in COP we divide the estimate of τ_i by the marginal utility of income α_i. Specifically, the permanence clause is associated to a switching cost of 24,079 COP (8.42 USD, approximately) on average per mobile.

7 Welfare analysis

To estimate the effect of permanence clauses on consumer, firm, and social welfare we compute four counterfactual scenarios. That is, two additional experiments with respect to the original study. Below we describe

[3]This is implied by computing: $\hat{\bar{\tau}} + 2 \times (\hat{\Sigma}_\tau + \hat{\Pi}_\tau) \approx -6.68 + 2 \times (1.16 + 1.25) < 0$.

the experiments computed for this appendix:

1. Eliminate permanence clauses before July, 2014 without price adjustment: Assume permanence clauses were nonexistent from the beginning of the sample period, hence the permanence clause indicator will take the value of zero, and assume that firms charge the same prices as in the observed equilibria.

2. Eliminate permanence clauses before July, 2014 allowing price adjustment: Assume permanence clauses were nonexistent from the beginning of the sample period, hence the permanence clause indicator will take the value of zero, and assume that firms charge the prices that result from the Betrand-Nash equilibria computed for every market.

3. Allow permanence clauses to exist across the whole sample period without price adjustment: Assume Resolution 4444 was never enacted, thus permanence clauses would still be in force, and assume that firms charge the same prices as in the observed equilibria.

4. Allow permanence clauses to exist across the whole sample period allowing price adjustment: Assume Resolution 4444 was never enacted, thus permanence clauses would still be in force, and assume that firms charge the prices that result from the Betrand-Nash equilibria computed for every market.

Table 2: Aggregate variation in consumer, firm, and social welfare due permanence clauses

Scenario	Δ C	Δ F	Δ S
Counterfactual 1			
Nonexistent permanence clauses and fixed prices	47,375	-15,348	32,027
Counterfactual 2			
Nonexistent permanence clauses and equilibrium prices	54,573	7,396	61,969
Counterfactual 3			
Permanence clauses during all sample and fixed prices	-201,604	20,942	-180,662
Counterfactual 4			
Permanence clauses during all sample and equilibrium prices	-185,065	-4,508	-189,573

Note: This table shows the aggregate variations in consumer, firm and social welfare implied by four counterfactual scenarios. Numbers are reported in millions of COP of June, 2016. Authors' calculations based on the data provided by GfK.

Table 2 presents the estimation of equations (19), (20), and (21) for each counterfactual scenario. All numbers are reported in millions of COP of June, 2016. As in the original paper, for the estimation of ΔF we multiply variations in firm surplus by the inverse proportion of units reported in our data to units sold in the national market[4]. That is, if η_t is the representativeness of our data during month t, then firm surplus is extrapolated to total market sales by multiplying units sold in the observed scenario and the counterfactual by $1/\eta_t$. Hence, we are assuming the unobserved portion of the market has the same sales composition as the one we observe. Given that variations are computed as the difference between the measure of welfare in the observed scenario and the measure of welfare in the counterfactual, a negative variation suggests the counterfactual situation generates higher welfare, while a positive variation means the observed scenario generates higher welfare.

In the first counterfactual, the elimination of permanence clauses from January 2014 would have increased overall welfare relative to the observed scenario by 32,027 million COP (11.24 million USD, approximately). This is explained by an increase in consumer welfare of 47,375 million COP (16.62 million USD, approximately) which compensates a decrease in firm profits by 15,348 million COP (5.39 million USD, approximately). For producers, such loses are particularly derived from the fact that the priced charged do not necessarily correspond to those that maximize individual profits.

In the second counterfactual, the elimination of permanence clauses from January 2014, and allowing firms to charge optimal prices would have increased overall welfare relative to the observed scenario by 61,969 million COP (21,73 million USD, approximately). This is explained both by an increase in consumer welfare of 54,573 million COP (19.15 million USD, approximately) and an increase in firm profits of 7,396 million COP (2.60 million USD, approximately). For producers, such gains are derived from more units sold compared to the observed scenario rather than from higher markups.

In the third counterfactual, allowing permanence clauses would have decreased social welfare by 180,662 million COP (63.40 million USD, approximately) relative to the observed scenario during the 23 months from July 2014 to June 2016. Losses for consumers would have been 201,604 million COP (70.74 million USD, approximately) whiÃśe gains for firms would have been 20,942 million COP (7.35 million USD, approximately). The positive variations in producer surplus are explained the higher prices charged by firms despite the presence of permanence clauses which should *a priori* be associated with subsidized prices for the sold handsets.

In the forth counterfactual, allowing permanence clauses and equilibrium prices would have decreased social welfare by 189,573 million COP (66.62 million USD, approximately) relative to the observed scenario during the 23 months from July 2014 to June 2017. Losses for consumers would have been 185,065 million COP (64.94 million USD, approximately) and for firms 4,508 million COP (1.58 million USD, approximately). The positive variations in producer surplus, which means lower profits from the counterfactual, are

[4]This information was also provided by GfK as mobile sales in Colombia per month.

explained by the fact higher markups do not outweigh lower sales in the counterfactual.

All counterfactual scenarios, specially the second and forth, show that the Colombian market for mobile terminals has been better off without permanence clauses. In the case of firms, we show network operators have been the ones experiencing greater benefits from the banning, despite their allegations of profit reduction and investment disincentives. In the case of consumers, we show that although there are some who derive higher utility from bundled sales with permanence clauses, their gains are significantly outweighed by the mass of consumers who experience lower utility levels from being locked-in.

8 Conclusions

In this appendix we discuss the results of a robustness check to the results presented in the study "Welfare effects of switching barriers through permanence clauses: Evidence from the mobiles market in Colombia". In particular, we measure the impact of Resolution 4444 enacted by the Communications Regulation Commission, which eliminated permanence clauses after July 2014. The variations with respect to the original study are mainly the addition of new data with an extra year of information and a more robust estimation of the parameters in which we focus only in estimating the demand side of the model.

Results show that switching costs both reduce the consumer's average utility and increase the variance of the utility distribution. In our counterfactual scenarios we compute the aggregate variations in consumer, firm and social welfare assuming for things: first, that permanence clauses were nonexistent in the Colombian market throughout the time series and that firms charged the same prices as in the observed data; second, that permanence clauses were nonexistent in the Colombian market throughout the time series and that firms charged the optimal prices suggested by a Betrand-Nash equilibrium; third, that permanence clauses were always in force with that firms charging the same prices as in the observed data; and forth, that permanence clauses were always in force with that firms charging the optimal prices suggested by a Betrand-Nash equilibrium. Results from the counterfactual scenarios also supported the hypothesis of permanence clauses being detrimental for consumer welfare, but in addition revealed they benefit firms basically because benefits experienced by network operators who sell bundled terminals and plans outweigh losses experienced by deterred competitors who would have sold terminals in cash price (with or without a calling option).

Our findings in this appendix serve as additional evidence supporting the conclusions of the main study, which contribute to the discussions of the effects of switching costs and bundled sales on social welfare and provides empirical evidence of the effects of bundle sales for the Colombian case and for markets with similar structure.

References

Berry, S., Levinsohn, J., and Pakes, A. (1995). Automobile prices in market equilibrium. *Econometrica*, 63(4):841–890.

Cameron, C. and Trivedi, P. (2005). *Microeconometrics: Methods and Applications*. Cambridge University Press.